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**Educational Expansion and its
Heterogeneous Returns for Wage Workers
– revised version –**

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Educational Expansion and its Heterogeneous Returns for Wage Workers

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Abstract

This paper examines the evolution of returns to education in the West German labour market over the last two decades. During this period, graduates from the period of educational expansion entered the labour market and an upgrading of the skill structure took place. In order to tackle the issues of endogeneity of schooling and its heterogeneous returns we apply two estimation methods: Wooldridge's (2004) approach that relies on conditional mean independence and Garen's (1984) control function approach that requires an exclusion restriction. For the population of workers from the SOEP, we find that both approaches produce estimates of average returns to education that decrease until the late 1990s and increase afterwards. The gender gap in returns to education seems to vanish. Furthermore, we find that the so-called "baby boomer" cohort has the lowest average return to education in young ages. However, this effect disappears over time.

JEL-classification: J21, J24, J31

Keywords: Returns to Education, Human Capital, Skill Upgrading, Wage Work

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1. Introduction

In Germany, a major expansion of higher secondary and tertiary education occurred during the 1960s and 70s. For instance, in the year 2005, 40 percent of the Germans in the age group of 25 to 30 held an upper secondary degree, as compared to 20 percent of the 55 to 60 year old (Statistisches Bundesamt 2006). A similar picture appears at the tertiary level, where 17% in the age group of 30 to 35 years held a higher technical college or university degree, compared to 13% of the 55 to 60 year old. Even though this was a moderate expansion from an international perspective (Müller and Wolbers 2003), it shares common goals. The expansion was issued to enhance individual well-being and equality of educational opportunity, among others. To some degree the latter goal seems to have been achieved (Müller and Haun 1994; Henz and Maas 1995; Schimpl-Neimanns 2000). However, the picture concerning the longer run labour market consequences of educational expansion is less clear. The reason is that when the students from the period of educational upgrading entered the labour market later on, this significantly raised the qualification structure of the workforce.

In our paper we investigate the question, whether the upgrading of schooling “devaluated” average long-run labour market returns. Our contribution is threefold. First, we investigate the evolution of heterogeneous returns to education in the twenty-year period from 1984 to 2006 to study the longer run labour market impacts of educational expansion. The empirical assessment is based on data from the German Socio-Economic Panel (SOEP). Second, we take the endogeneity and selectivity of school choice (see Card 1999, Willis and Rosen 1979, among others) into account, which results in models of heterogeneous rates of returns (see Blundell, Dearden and Sianesi 2005, Flossmann and Pohlmeier 2006, among others and Heckman, Lochner and Todd 2008 for a discussion on the economics of rates of returns studies). A correlated random coefficient model is employed, where the explanatory variable “years of schooling” is measured as a continuous treatment variable, which can be correlated with unobserved heterogeneity.

Identification is based on different assumptions. Following Wooldridge (2004) we identify the average return to education via conditional mean independence assumptions and following Garen (1984) a control function approach is employed which uses exclusion restrictions to control for selection on unobservable heterogeneity. Third, the returns to education in West Germany are differentiated by gender and birth cohort. In this way, we investigate the relationship induced by female labour force participation and the rise of newborns until the 1960s and its decline afterwards.

Our findings indicate that the average returns to education decreased until the late 1990s and increased afterwards. Using Wooldridge's approach, our results vary between 4.9 and 6.6 percent for the average partial effect of an additional year of schooling, which seems to be at the lower end of previous findings for Germany (Boockman and Steiner 2006, Lauer and Steiner 2001, Flossmann and Pohlmeier 2006, among others). Regarding the gender aspects, the average returns to education seem to have been larger for women during the 1980s and early 1990s. However, the gap decreases over time, which may be a consequence of increased participation of women. Furthermore, we find that the so-called "baby boomer" cohort has the lowest average return to education compared to the cohort before and the one thereafter (the former is characterised by lower and the latter by sharply decreasing cohort sizes). While this finding seems to be in line with the literature on wages and cohort size (Macunovich 1999), in our data the effect seems to vanish with age.

This paper is organized as follows. Section 2 discusses factors that influence returns to education over time. In section 3, we develop the idea of heterogeneous returns to education in a correlated random coefficient model, and compare estimates from the conventional, as well as Wooldridge's (2004) and Garen's (1984) approaches. Section 4 describes the data set and variables used. Furthermore, first descriptive results for the evolution of educational attainment over time are presented. In section 5, we discuss estimation results differentiated by estimation techniques, gender and cohorts over time. Section 6 concludes.

2. Educational Expansion, Wages and the Labour Market in West-Germany

Educational attainment started to increase in the 1960s in Germany leading with a lag to the upgrading of educational qualifications in the labour market (Bock and Timmermann 1998). In our sample of workers from West-Germany, extracted from the SOEP, average years of schooling increased for women (men) from 10.8 (11.6) years in 1984 to 12.4 (12.7) years in 2006. In recent years male and female average educational attainment became similar. In a standard economic supply and demand labour market framework, a rising supply of (high-) skilled workers induces, *ceteris paribus*, a decline in the returns to education. A related concern is that educational expansion may have resulted in institutions starting to accept students increasingly from the lower end of the distribution of student abilities so that weaker students might have been admitted to higher education, leading to a decrease in the average productivity level of higher educated workers.

However, besides the educational expansion, there exist other factors that have influenced demand and supply conditions on German labour markets over the last two decades. Some important factors have been, for instance, increasing female labour market participation, birth cohort sizes, wage determination processes, and skill-biased technological change. In West-Germany, the female participation rate has been rising during the last decades, leading to a catching-up to men and competition for college slots and labour market positions. Based on the decreasing gender-gap in educational attainment and labour market participation, a convergence of gender-specific returns to education can be expected.

West-Germany, as well as many other western countries, experienced a demographic change due to a baby boom that peaked during the mid-1960s and sharply decreasing cohort sizes afterwards. Changes in the number of births alter the supply of workers entering the market about twenty years later, i.e. during our period of investigation. If larger birth cohorts enter the labour market and substitution in production is limited between younger and older workers, *ceteris paribus*, a downward pressure on returns to education for labour market entrants arises (Macunovich 1999; Freeman 1979, among others). Therefore, one may expect decreasing returns

to education for the baby boom cohorts and increasing returns to education for individuals born after 1964 when cohort sizes started to decline sharply. In addition, there was fierce wage competition for entrants due to unemployment rates as high as ten percent in Germany. Compared to entrants, incumbent workers in Germany enjoy some protection against wage competition due to strong unions and/or efficiency wage considerations (Franz and Pfeiffer 2006). Because large cohorts are absorbed gradually by the labour market when experience increases, we expect lower returns to education for labour market entrants.

The computer revolution that started around 1970 changed the organisation of labour away from routine manual tasks to non-routine analytical and creative tasks (Autor, Katz and Kearney 2006; Spitz-Oener 2006, among others). The demand shift towards analytical skills presumably favoured the high skilled and may even have increased returns to education, despite increasing supply (Acemoglu 2002, among others).

To sum up, we expect supply side factors like educational expansion and the increase in female participation to lower the returns to education (in a *ceteris paribus* sense). Similarly, supply side factors such as a decreasing cohort size and demand side factors such as skill-biased technological change and workplace innovations are likely to increase the returns to education. In Germany, the impact of educational expansion on wages is also likely to be formed by the process of wage determination, the regulation of labour as well as the rate of unemployment and active labour market policies. We would like to analyze the empirical evolution of the returns to education in West Germany from 1984 to 2006 that resulted from the factors outlined above.

3. Econometric Approach

Our empirical framework is the correlated random coefficient model (Heckman and Vytlacil 2001, see also Blundell et al. 2005; Björklund and Kjellström 2002, Heckman et al. 2008):

$$\ln Y_i = a_i + b_i S_i \text{ with } a_i = a'X_i + \varepsilon_{ai} \text{ and } b_i = b'X_i + \varepsilon_{bi} \quad (1)$$

where the outcome variable, $\ln Y_i$, is the natural log of wages and the explanatory variable, S_i , is years of schooling of individual i . This equation has been derived from optimal schooling choice where education is determined by a respective individual's observed and unobserved marginal benefits and the costs of schooling (Card 1999). The model has an individual-specific intercept a_i and slope b_i that may depend on observable variables X_i and unobservable heterogeneity ε_{ai} and ε_{bi} . The heterogeneity components capture influences from gender, family background, age, preferences, ability, etc. such that a_i and b_i represent random coefficients. We do not assume that b_i and S_i are independent. a_i and S_i as well as b_i and S_i can be correlated (Wooldridge, 2004). Since individuals with higher expected benefits from education are more likely to participate longer in education, the returns to education b_i may in general be correlated with S_i if variation in unobserved (to the econometrician) benefits implies positive self-selection. In this case, the schooling variable is influenced by its own coefficient, yielding an endogeneity problem.

Our research interest is the effect of S_i on $\ln Y_i$ represented by b_i in equation (1). In this model, the return to education varies across individuals in both, the observable heterogeneity in returns X_i and the unobserved individual-specific returns to schooling, ε_{bi} . The resulting distribution of returns will be summarized with the average partial effect, APE (equation 2, Flossmann and Pohlmeier 2006; Wooldridge 2004). APE measures the average return per additional year of education for a randomly chosen individual from our population:

$$E(\partial \ln Y / \partial S) = E(b_i) = \beta \quad (2)$$

The earnings equation (1) nests more specific models. If returns to education are homogenous, the outcome equation can be re-written as the classical Mincer-type of earnings function (Blundell and Costas 2000):

$$\ln Y_i = a'X_i + \bar{b}S_i + \varepsilon_{ai} \quad (3)$$

where \bar{b} is the common rate of return. Unobserved heterogeneity may exclusively enter the intercept of the wage equation but not the slope coefficient. In that case there might still be endogeneity problems, if the unobserved general individual earnings capacity ε_{ai} is correlated with S_i . One appealing feature of the general approach (1) is that variation in unobserved heterogeneity affects the slope as well, i.e. that unobserved heterogeneity influences the wage effect of education.

3.1. *Conventional Methods*

When estimating (1) by OLS, there are three potential sources of bias. First, if individuals with high absolute earnings capacity both acquire more education and earn higher wages, schooling S_i will be positively correlated with ε_{ai} (Griliches 1977). This ability bias induces an upward bias in the estimated average return (Behrman and Rosenzweig 1999). Second, classical measurement error in the schooling variable S_i induces a downward bias (Griliches 1977). Third, there can be a bias if individuals differ in their relative earnings capacity and act upon their comparative advantage when choosing their level of education (Willis and Rosen 1979). If returns to education are homogeneous, the latter bias is absent.

In the literature on the return to schooling in Germany instrumental variables (IV) methods are common to handle the endogeneity problems. For instance, Lauer and Steiner (2001) estimate homogeneous returns to education using different family background variables as instruments. The results depend on the instruments used and vary between 6.6 and 14.8 percent. However, when schooling is also correlated with unobserved individual heterogeneity, standard IV methods may fail to identify APE. Heckman and Li (2004). Ichino and Winter-Ebmer (1998), among others instead estimate the local average treatment effect (LATE) of schooling for Ger-

many using different instruments. Since each instrument implies its own LATE and the group of compliers cannot be identified without further assumptions, this may be regarded as a drawback. However, LATE is especially interesting when school reforms are used as instruments since LATE measures the returns to schooling for those who changed their level of schooling because of the reform. With this approach, Pischke and van Wachter (2008) find rather low marginal returns to education in Germany. In our empirical analysis, we employ methods that reduce the potential bias from OLS and IV techniques. Furthermore, in contrast to the LATE interpretation of IV techniques, we are interested in assessing the APE.

3.2. Wooldridge's (2004) Conditional Mean Independence (CMI) Approach

The methods rely on different identifying assumptions: Wooldridge's (2004) conditional mean independence (CMI) approach and Garen's (1984) control function (CF) approach. According to Wooldridge (2004) APE is identified by the following two assumptions if the linear outcome equation (1) holds:

$$E(\ln Y_i | a_i, b_i, S_i, X_i) = E(\ln Y_i | a_i, b_i, S_i) = a_i + b_i S_i \quad (4)$$

$$E(S_i | a_i, b_i, X_i) = E(S_i | X_i) \text{ and } Var(S_i | a_i, b_i, X_i) = Var(S_i | X_i) \quad (5)$$

where the elements of X_i are suitable proxy variables for the observed and unobserved heterogeneity, i.e. the X_i should be good enough predictors of S_i . The first assumption postulates that the vector X_i is redundant given S_i and (a_i, b_i) in the structural conditional expectation (4). This identification assumption obviously holds since the control variables X_i enter the earnings function through a_i , b_i , and S_i only. The second assumption is a redundancy condition of the form that both heterogeneity terms a_i and b_i are redundant in the first two conditional moments of the schooling variable S_i conditional on a set of covariates X_i . The latter is the strongest assumption as it requires a differentiated set of variables that control sufficiently for observable and unob-

servable heterogeneity. These conditional moment independence (CMI) conditions are a weaker form of conditional independence assumptions (CIA) (Wooldridge, 2002: 607). Based on assumptions (4) and (5) Wooldridge (2004) derives the following estimator for APE:

$$\hat{\beta} = \frac{1}{N} \sum_{i=1}^N \left(\left(S_i - \hat{E}(S_i|X_i) \right) \ln Y_i \right) / \hat{Var}(S_i|X_i) \quad (6)$$

Because $\ln Y_i$, S_i and X_i are observable, one needs to estimate the conditional mean and variance, $E(S_i|X_i)$ and $Var(S_i|X_i)$. Since S_i is nonnegative, simple linear models have shortcomings. Therefore, we employ a generalized linear model (GLM) with a Poisson distributional assumption for years of schooling S_i :

$$E(S_i|X_i) = e^{\gamma'X_i} \text{ and } Var(S_i|X_i) = \sigma^2 e^{\gamma'X_i} \quad (7)$$

This specification guarantees positive estimates of both conditional mean and variance. Contrary to the standard variance assumption of equality between the conditional variance and the mean equation, (7) relies on the weaker Poisson GLM variance assumption that allows the variance-mean ratio to be any positive constant (Wooldridge, 2002). A consistent estimator of σ^2 is obtained as the mean of squared Pearson residuals. Since analytical standard errors have not been developed so far, standard errors of the APE are bootstrapped.

3.3. Garen's (1984) Control Function (CF) Approach

Garen (1984) proposed a possible alternative solution to the random coefficient estimation problem - called the control function (CF) approach - that is similar to Heckman's (1978) two-step estimator. While the standard IV approach does generally not identify APE in the heterogeneous returns context the CF approach does. The CF approach is implemented by a simul-

taneous modelling of both schooling and wages. Hence, an explicit model of the schooling selection process, which relates the rule for assigning individuals to treatment to the potential treatment outcomes is required:

$$S_i = c'X_i + dZ_i + v_i \text{ with } E(v_i|Z_i, X_i) = 0 \quad (8)$$

where both X_i and Z_i influence the educational decision and v_i represents the usual error, incorporating unobserved components which determine the choice of education. Z_i is an exclusion restriction, i.e. it should have no correlations with unobserved heterogeneity in the wage equation. The error terms v_i , ε_{ai} and ε_{bi} are normally distributed with zero means and positive variances that are possibly correlated with each other.¹ Following Garen (1984) one can formulate an augmented wage equation of the form:

$$\ln Y_i = a_i + \beta S_i + \gamma_1 v_i + \gamma_2 v_i S_i + w_i \quad (9)$$

where $\gamma_1 v_i$ and $\gamma_2 v_i S_i$ are the control functions with $\gamma_1 = \text{cov}(\varepsilon_{ai}, v_i) / \text{var}(v_i)$ and $\gamma_2 = \text{cov}(\varepsilon_{bi}, v_i) / \text{var}(v_i)$. Once these terms are included in the outcome equation (and implicitly subtracted from its error term), the error term w_i has all the desirable properties, i.e. it is orthogonal to all of the regressors in the new equation: $E(w_i|X_i, S_i, v_i) = 0$ (Heckman & Robb, 1985).

This model can be estimated using a generalization of the two-step approach. In the first step an estimation of the schooling choice is used to construct the control functions that are included as additional regressors in the augmented wage equation. The estimated coefficients of v_i

¹ This trivariate normality assumption can be weakened to the condition that conditional expectations of the unobserved earnings components ε_{ai} and ε_{bi} are linear in the residual of the selection equation (Blundell et al., 2005).

and $v_i S_i$ provide information about the selection on the unobserved absolute earnings capacity term and about selection on the comparative earnings capacity, respectively. If an individual attains a higher (lower) level of education than according to our expectations, the value of v_i is positive (negative). For example, if the coefficient γ_1 of the first control function is positive, this implies that the unobserved factors that lead to educational ‘over-achievement’ (positive v_i) have a positive impact on earnings. The sign of the coefficient γ_2 of the second control function describes how this effect changes with increasing levels of education. Following the comparative advantage hypothesis (Willis and Rosen 1979), we expect that γ_2 is positive, i.e., those with unexpectedly large amounts of schooling (positive v_i) tend to earn more than the others with higher education. Based on their higher unobserved marginal returns they select into higher education according to their comparative advantage.

4. Data and Descriptive Analysis

The empirical analysis is based upon samples from 23 waves of the German Socio-Economic Panel Study (SOEP, see Haisken-DeNew and Frick 2005). SOEP contains information on education, employment and earnings as well as retrospective information about family background. We include SOEP refreshment samples from 1998 and 2000. Due to lack of comparability, foreign-born individuals were excluded from the sample. Furthermore, the analysis is restricted to West-German citizens for comparison reasons. Self-employed workers are excluded from the sample since they are exposed to different earnings-generating mechanisms. The resulting sample is composed of full-time dependent workers aged between 25 and 60 that work 30 hours or more per week. After eliminating observations with missing values we obtain a final sample size that ranges from 1,535 observations in 1984 to 3,965 in 2000.²

The dependent variable is the natural logarithm of real earnings per hour worked. The measure of years of schooling is derived by attaching a standard number of years to the highest

² Although sample size almost doubles due to refreshment samples, the following estimation results do not change substantially if the new samples are excluded.

educational level (cp. table 1). As control variables, gender and individual's age (in linear and quadratic terms) are included. We use age variables instead of potential labour market experience because the latter might be endogenous with respect to schooling.

- Table 1 about here -

To justify the conditional moment independence assumptions a set of family background information is utilized that is covered by recall questions and that is available for a sufficient number of persons in each wave considered. Family background and parental educational and occupational attainment proxies the parental influence on educational attainment and later employment carriers (Erikson and Jonsson 1996, among others). The measure for parents' education follows the CASMIN educational classification, which has the advantage to combine information on the highest school degree and the highest vocational degree of the parents (Erikson and Goldthorpe 1992). The CASMIN categories have been summarized in five categories for fathers and in a dummy-variable for mother's higher education (cp. Table 1).

There are four categories of parents' occupational position (cp. Table 1). These categories should proxy for the economic circumstances of the family, which affect educational choice by influencing costs of schooling. A further proxy for costs of schooling is the dummy variable "rural socialisation" (see Card 1995 among others).

For the control function approach the number of siblings is used as an exclusion restriction. We assume that it satisfies the two conditions for valid instrumental variables (Wooldridge, 2002). First, Becker and Tomes (1976) and Hanushek (1992), among others, hypothesize a positive correlation between the number of siblings and individual educational attainment even after controlling for other family background characteristics. Parents try to optimally allocate financial and non-financial resources to their children who compete for the attention and resources of their

parents. Therefore, educational achievement and total family size might be negatively related given limited educational resources.

Second, the instrumental variable should be uncorrelated with unobserved individual's earnings capacities, i.e., the number of siblings should not have an effect on income other than the indirect effect transmitted over educational attainment. Because we control for a set of other family background variables like parents' education, occupation and the place of socialisation we do not expect a non-negligible, systematic and independent effect of the number of siblings on earnings. In the case of Wooldridge's (2004) CMI approach the number of siblings serves as a further control variable. Table 1 gives an overview of the variables and its definitions. Table 2 provides summary statistics for key individual level variables in selected years.

- Table 2 about here -

5. Estimation Results

5.1. Evolution over Time: Comparison of Different Estimation Techniques

Figure 1 compares the evolution of our three different estimates of individual returns to education in West Germany during the period 1984 to 2006. Besides, the results from a Mincerian OLS regression with years of schooling and controlling for age in linear and squared functional form on log wage, the APE from the conditional mean independence (CMI) approach and from the control function (CF) approach are graphed. With OLS we find a slight downward trend in the evolution of returns to schooling until the late 1990s. The returns to one additional year of education fell from 7.4 percent in 1984 to 5.4 percent in 1998. From 1998 onwards, we find increasing returns to education reaching a new local maximum of 6.8 percent in 2002. The estimates until 1998 are in line with the findings of Lauer and Steiner (2001), among others. We are not aware that the increase in returns starting around 1998 has been documented so far.

- Figure 1 about here -

Interestingly, the APE estimated under conditional moment independence (CMI) follows a fairly similar evolution pattern over time. Both approaches produce also similar small standard errors varying between 0.003 and 0.004 in the observation period. However, there are differences. First, the estimated APE is always lower than standard OLS, between 0.2 and 1 percentage points. According to our interpretation this difference reflects the potential ability bias from OLS estimates. Taking into account the heterogeneity of returns to education and controlling for family background variables the CMI approach controls to a certain degree for positive ability. Second, although OLS and APE estimates are comparable over time, their content varies. APE measures the average of the distribution of heterogeneous returns, whereas OLS measures the average return to education, i.e. that is homogenous for all individuals. Compared to the literature, our estimate of the APE seems to be rather low. Maier, Pfeiffer, and Pohlmeier (2004), for instance report an estimated APE of 8.7 percent for the year 1999 for German male workers.

The CF approach has been implemented in a two-stage estimation procedure (for detailed estimation results see Table 3). The first stage, the educational attainment selection equation, has been used for testing the validity of the instrumental variables. A regression that includes the number of siblings in a simple OLS log-wage equation together with other family background variables was insignificant suggesting that the number of siblings seems to be a reasonable exclusion restriction. According to our findings the number of siblings has a strongly negative influence on educational attainment, holding constant other family background characteristics. This seems to be in line with the literature mentioned in the previous section above.

- Table 3 about here -

The coefficient of the control function for the selection on unobserved absolute earnings capacity is positive, although it decreases over time (cp. Table 3) and is never significant. Thus, we find no evidence for a positive ability bias with the control function approach in our data. Educational background of the family proxies the absolute earnings capacity. The coefficient of the control function for the selection on comparative earnings capacity is always negative, which contradicts the comparative advantage hypothesis. In our data we find that those with unexpectedly high amounts of schooling have lower marginal returns to education, which seems to be similar to findings from Maier et al (2004) and Pischke and van Wachter (2008). There are individuals, who have done worse after more schooling. The effect is significant in the period from 1984 to 1989 and from 2000 onwards. In these years, we find negative selection on unobservable returns.

The evolution pattern of the estimated APE under CF approach deviates substantially from the CMI results. First, the yearly estimates derived from the CF approach are more volatile and less precise. Hence, the standard errors are higher.³ Detailed tests show that the deviations are usually not significant during the 1980s and the initial fluctuations (see Figure 3) might be a consequence of lesser observations for the 1980s. Second, there is a stronger and significant increase of the APE after 1998 compared to the CMI approach. In 2006, the APE is 13 percent, which is in line with some recent IV studies for Germany, Flossmann and Pohlmeier (2006). From a methodological point of view the deviations reflect differences in the identification strategies. From a substantive point of view the discrepancy during the last years could be a hint for rising selection on unobservables, which the CF approach controls for (Taber, 2001). This coincides with the significant effects for the second control function from 2000 onwards.

To summarize our findings so far: Independently of the method used, the returns to education (APE) were fairly constant during the 1980s and 1990s in (West-) Germany. Despite a continuous upgrading of educational qualification however, they started to increase from 1998

³ They vary around 0.04 in the 1980s and between 0.02 and 0.03 afterwards, i.e. they are about ten times larger than the CMI and OLS standard errors.

onwards, although only moderately. This finding seems to be line with rising wage inequality in Germany that set in around 1994 (see Gernandt and Pfeiffer 2007).

5.2. Gender and Cohort Effects

The following analysis differentiates gender and birth cohorts to take a second look at the recent increase in the APE. The comparison rests on the CMI approach because it produces lower standard errors than the CF approach. The average return to an additional year of education in our population of women from the SOEP declined from 8.6 percent in 1984 to about 4.9 percent in 1996 (cp. Figure 2). According to our interpretation this is mainly the result of the female educational expansion and rising participation, a supply side interpretation, intensified by non-neutral technical progress (Spitz-Oener 2006) and/or gender concerns in collective wage bargaining. Interestingly, since 1999 the APE is increasing again for women, as it is for men. Furthermore, Figure 2 indicates that the gender gap in returns to education decreased over time, which is in line with findings from Lauer and Steiner (2001), among others. Tests indicate that the differences became insignificant after 1995.

- Figure 2 about here -

In a further step, the average returns to education are compared for four different birth cohorts. We differentiate the cohorts based on relative birth cohort sizes to avoid interpretation problems stemming from self-selection into the labour market. Cohort size in the labour market might be endogenous because individuals change their educational attainment and labour market entry with respect to cohort size (Berger, 1989; Macunovich, 1999). Each cohort is composed of eight years: people born between 1942 and 1949, people born between 1950 and 1957, people born between 1958 and 1965 and those born between 1966 and 1973. The cohort boundaries are geared to cohort sizes: The oldest cohort has low birth rates due to the 2nd World War and the

post-war period. The second cohort 1950-57 is of relatively constant size, whereas the third cohort 1958-65 is the “baby boom” cohort with strongly increasing cohort sizes peaking in 1964. Finally, the last cohort 1966-73 is characterized by a sharply declining cohort size. In order to have cohorts with a sufficient number of observations, our estimations are restricted to birth cohorts that are older than 27 to 34 years, e.g. we estimate APE for cohorts born 1958-65 starting at the year 1992.

Time, cohort and life cycle effects cannot be disentangled empirically because it is impossible to observe two different birth cohorts at the same age and in the same year (Heckman & Robb, 1985). To empirically assess cohort effects in average returns to education in Germany different cohorts at the same age are compared at different points in time in the same labour market. Both Figures 3 and 4 display estimation results for the four cohorts at a given age. We follow the development in the returns to education over a specific phase of the working life of a cohort for a period of five years. For example, all cohorts in Figure 3 are observed at ages 27-34. However, we do this for the cohort 1950-57 in 1984, for the cohort 1958-65 in 1992 and for the cohort 1966-73 in 2004.

- Figure 3 about here –

- Figure 4 about here -

Figure 3 reveals that the “baby-boomer” cohort has the lowest average return to education compared to the cohort before (1950-57) and the one thereafter (1966-73). A larger cohort size seems to reduce the average return to education at young ages (27-38 years old). Although these differences are not well determined from a statistical point of view, they weakly support that higher supply of labour market entrants increases wage competition, which seems to be also in line with findings from Boockmann and Steiner (2006) and Lauer and Steiner (2001), among others. As a new result, we find that the quantitative differences in the APE disappear when the

“baby-boomer” cohort is compared at older ages (35-46 years old) with other cohorts at the same age (see Figure 4). If so, cohort effects seem to exist only for the young when they enter the labour market and seem to vanish over time.

6. Conclusions

In Germany, graduates from the period of educational expansion in the 1960s and 70s entered the labour market during the period of observation from 1984 to 2006. With a lag, this educational expansion contributed to skill upgrading of the labour force. In our sample from the SOEP the average years of education increased by roughly one year in this period. In order to tackle the issue of endogeneity of school choice and its heterogeneous returns we applied two estimation methods: Wooldridge’s (2004) CMI approach and Garen’s (1984) CF approach. The former method relies crucially on the conditional moment independence assumption, which requires sufficient observable control variables. The latter method employs distributional assumptions and needs an exclusion restriction such that it can control for selection on unobservables.

Our empirical findings indicate that both approaches produce estimates of average returns to education that decrease until the late 1990s and increase afterwards. During the period from 1984 to 2006 the estimated APE follows a roughly similar evolution pattern over time although standard errors from Garen’s approach are larger. According to the Wooldridge approach, returns to one additional year of education fell from 6.5 percent in 1984 to 4.9 percent in 1998 and to 6.5 percent in 2002 again. During the 1980s and early 1990s returns to education have been higher for women than for men, but the gender gap in returns vanish after 2000. Furthermore, we find that the cohorts of “baby boomers” (workers born between 1958 and 1965) had the lowest average return to education. However, the effect exists only at young ages and disappears when employees become older.

In this study education is measured as years of schooling. Future research could be directed to specific characteristics of the German educational system, like early ability tracking

and dual vocational educational qualifications. In addition, research could be directed to other outcomes of educational expansion, like economic growth, innovativeness and unemployment.

7. References

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Figure Captions

Figure 1

Returns to Education, 1984-2006: OLS, CMI, and CF approach compared

Figure 2

Average Partial Effect (CMI approach) by gender, 1984-2006

Figure 3

Average Partial Effect (CMI approach) by birth cohorts at same age 27-38

Figure 4

Average Partial Effect (CMI approach) by birth cohorts at same age 35-46

Figure 1

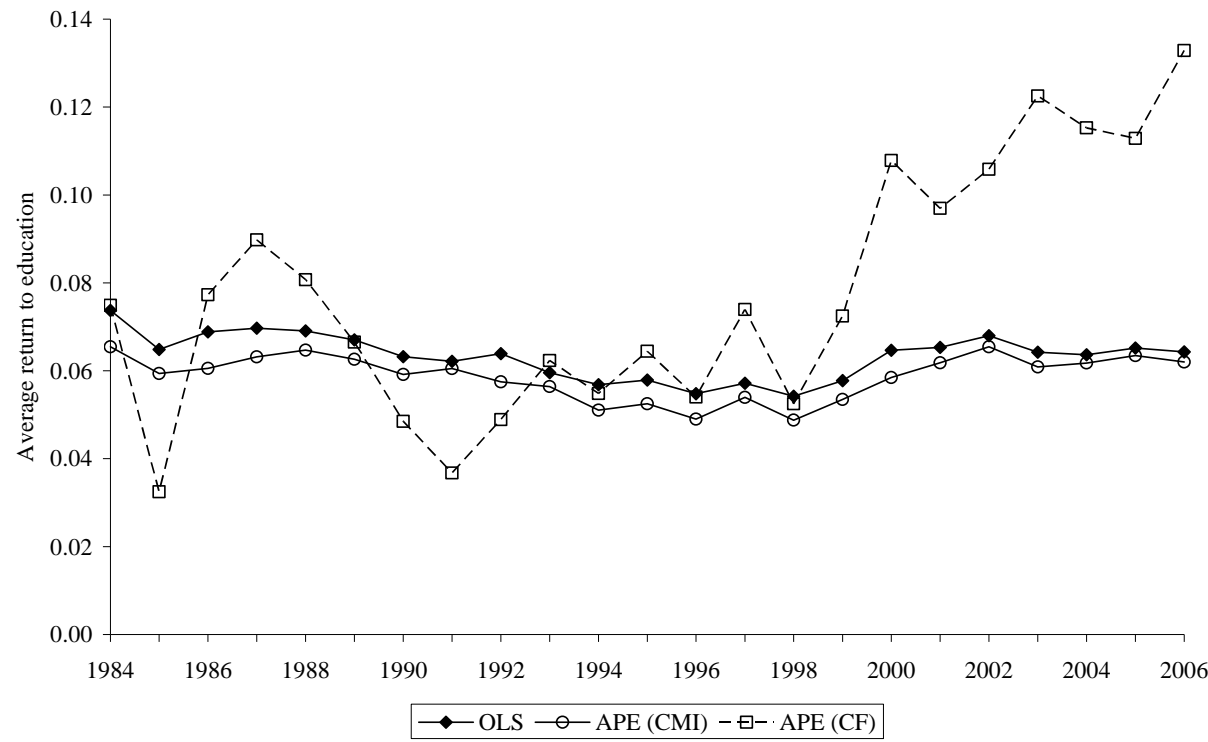


Figure 2

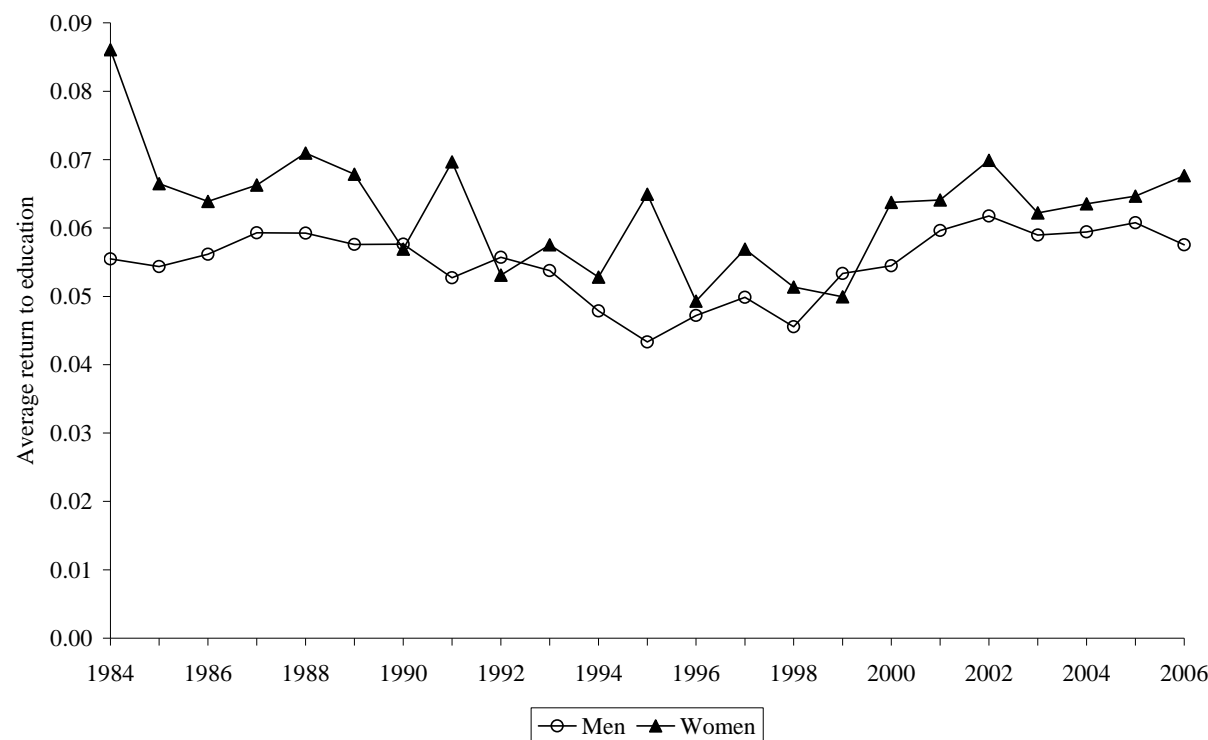


Figure 3

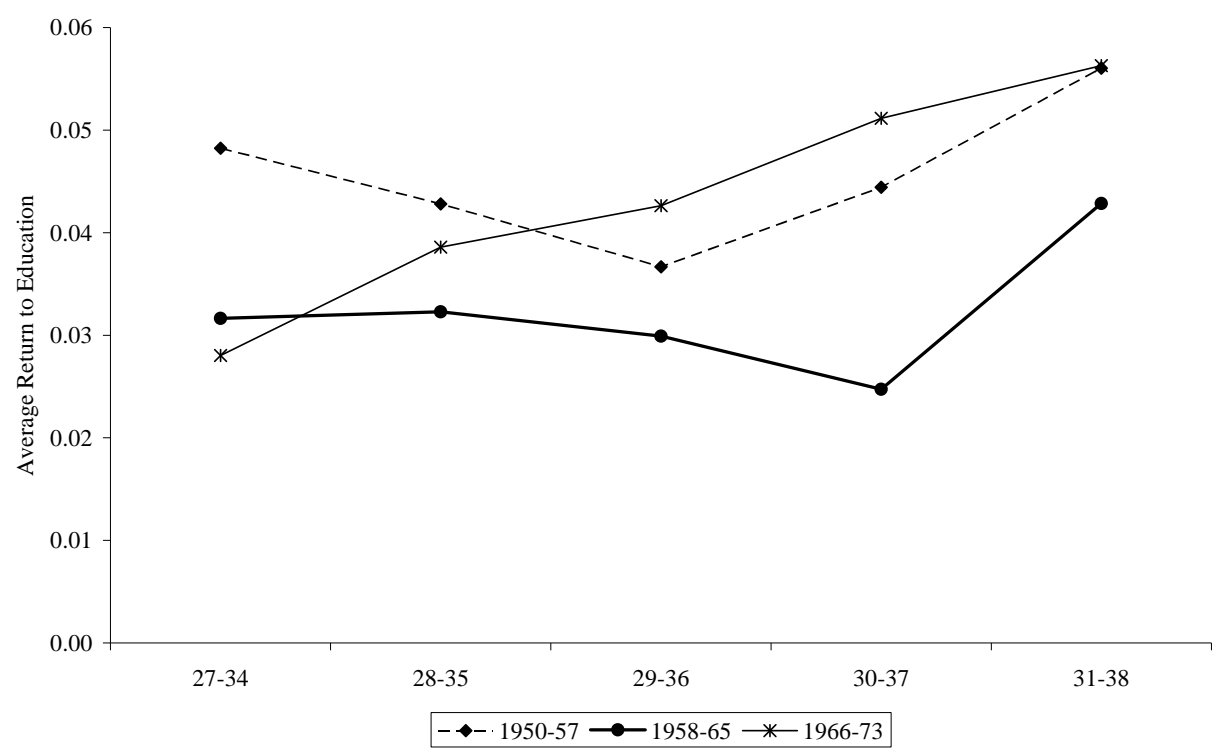


Figure 4

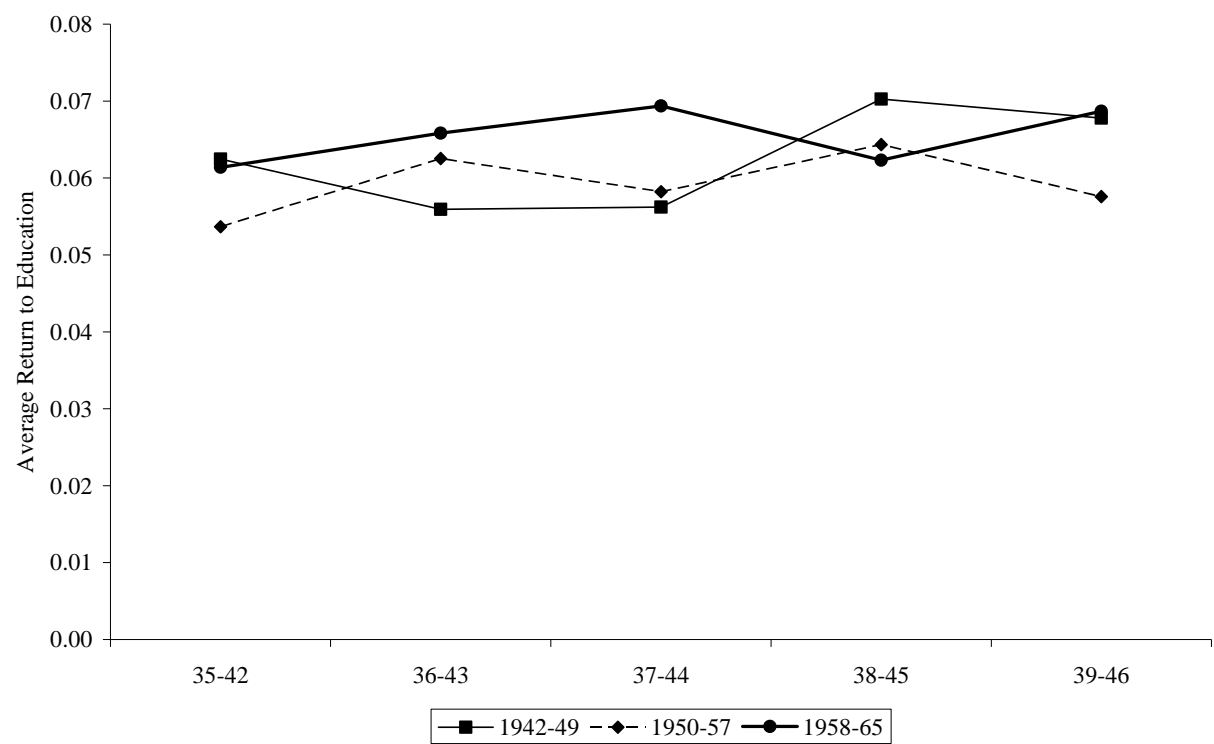


Table 1

Variable Definitions

<i>Variable Name</i>	<i>Description</i>
log wage	Log gross hourly wage
years of education	Years of education: constructed with standard times for highest educational and vocational degree obtained: no degree (7 years), lower secondary (9 years), intermediate secondary (10 years), technical secondary (12 years), higher secondary (13 years), vocational training (+1.5 years), vocational school (+2 years), higher technical college (+3 years), university (+5 years)
<i>Demographics</i>	
age	Age in years
age squared	Age squared
female	Dummy for sex (1= female; 0= male)
<i>Father's Education</i>	
father elementary educ	Reference category: Inadequately completed elementary education or (compulsory) elementary education
father compulsory and vocational educ	Compulsory education plus vocational training
father secondary intermediate educ	Secondary intermediate education, with/without vocational training
father full secondary educ	Full secondary education (Abitur), with/without vocational training
father university educ	University/ University of applied sciences
<i>Mother's Education</i>	
mother low educ	Reference category: Inadequately completed elementary education or (compulsory) elementary education; compulsory education plus vocational training
mother high educ	Secondary intermediate education, with/without vocational training Full secondary education (Abitur), with/without vocational training University/ University of applied sciences
<i>Occupational Position Father</i>	
father blue collar	Dummy (1= father blue collar; 0 else)
father white collar	Dummy (1= father white collar; 0 else)
father self-employed	Dummy (1= father self-employed; 0 else)
father civil servant	Dummy (1= father civil servant; 0 else)
<i>Place of Socialisation</i>	
rural socialisation	Dummy (1= rural socialisation, i.e. countryside; 0= urban socialisation, i.e. city, big town, small town)
<i>Family Composition</i>	
number siblings	Number of siblings

Table 2

Summary Statistics

	1984		1995		2006	
	Mean	Std. dev.	Mean	Std. dev.	Mean	Std. dev.
log wage	2.36	0.42	2.56	0.41	2.58	0.50
years of education	11.68	2.48	12.20	2.66	12.77	2.73
age	40.21	9.81	40.18	9.66	43.10	8.87
age squared	1,712.85	810.46	1,707.61	813.04	1,936.14	756.93
female	0.35	0.48	0.39	0.49	0.43	0.50
father elementary educ	0.17	0.37	0.12	0.33	0.12	0.32
father compulsory and vocational educ	0.65	0.48	0.64	0.48	0.61	0.49
father secondary intermediate educ	0.09	0.28	0.11	0.32	0.14	0.35
father full secondary educ	0.03	0.18	0.04	0.18	0.03	0.18
father university educ	0.07	0.25	0.09	0.28	0.10	0.30
mother high educ	0.12	0.33	0.17	0.38	0.23	0.42
father blue collar	0.49	0.50	0.47	0.50	0.45	0.50
father white collar	0.18	0.38	0.14	0.35	0.13	0.34
father self-employed	0.16	0.36	0.21	0.41	0.25	0.44
father civil servant	0.11	0.31	0.12	0.33	0.11	0.31
rural socialisation	1.70	1.78	1.68	1.68	1.66	1.64
number siblings	0.40	0.49	0.37	0.48	0.38	0.49
N	1,545		2,075		3,134	

Table 3

Returns to Education, 1984-2006: CF approach

	1. Stage		2. Stage				
	IV: Number of sib- lings		Selection on unobserved absolute earnings capacity		Selection on unobserved com- parative earnings capacity		N
	coeff.	(s.e.)	coeff. γ_1	(s.e.)	coeff. γ_2	(s.e.)	
1984	-0.091***	(0.032)	0.034	(0.081)	-0.003***	(0.002)	1,545
1985	-0.129***	(0.032)	0.081	(0.132)	-0.004**	(0.002)	1,600
1986	-0.094***	(0.031)	0.053	(0.096)	-0.005***	(0.002)	1,682
1987	-0.122***	(0.031)	0.062	(0.054)	-0.006***	(0.002)	1,775
1988	-0.132***	(0.031)	0.032	(0.047)	-0.003***	(0.002)	1,798
1989	-0.128***	(0.030)	0.059	(0.047)	-0.004***	(0.002)	1,922
1990	-0.151***	(0.029)	0.036	(0.035)	-0.002	(0.001)	2,007
1991	-0.148***	(0.028)	0.050	(0.033)	-0.002	(0.001)	2,122
1992	-0.164***	(0.028)	0.030	(0.031)	-0.002	(0.001)	2,107
1993	-0.180***	(0.028)	0.032	(0.036)	-0.003*	(0.002)	2,124
1994	-0.185***	(0.029)	-0.004	(0.031)	0.000	(0.002)	2,082
1995	-0.188***	(0.031)	0.023	(0.032)	-0.002	(0.002)	2,075
1996	-0.182***	(0.031)	0.024	(0.033)	-0.002	(0.002)	2,057
1997	-0.190***	(0.031)	0.018	(0.032)	-0.003*	(0.002)	2,011
1998	-0.220***	(0.031)	0.009	(0.026)	-0.001	(0.001)	2,145
1999	-0.211***	(0.031)	0.001	(0.029)	-0.001	(0.001)	2,163
2000	-0.158***	(0.023)	0.011	(0.031)	-0.004***	(0.001)	3,965
2001	-0.154***	(0.023)	-0.001	(0.028)	-0.002***	(0.001)	3,961
2002	-0.137***	(0.024)	0.009	(0.035)	-0.003***	(0.001)	3,668
2003	-0.158***	(0.025)	-0.010	(0.035)	-0.004***	(0.001)	3,476
2004	-0.150***	(0.025)	-0.011	(0.036)	-0.003***	(0.001)	3,366
2005	-0.153***	(0.026)	-0.010	(0.040)	-0.003**	(0.002)	3,220
2006	-0.157***	(0.027)	-0.001	(0.038)	-0.005***	(0.001)	3,134

Notes: (1) The first stage includes additional regressors such as gender, age, rural socialisation, educational and occupational background of the parents. (2) The second stage includes additional regressors such as years of education, gender, age, rural socialisation, educational and occupational background of the parents. The IV number of siblings is excluded. (3) Standard errors on the second stage are bootstrapped each with 500 repetitions. (4) Significant: *** at the 1% level; ** at the 5% level; * at the 10% level.